

Composition Bias and Italian Wage Rigidities over the Business Cycle*

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Abstract

I estimate the cyclical behavior of Italian real wages over the period 1985-2003 controlling for the so-called “composition bias”. Aggregate real wage statistics, commonly used to measure real wage elasticity, are affected by the bias arising from the cyclical change in the skill-composition of the labor force. An analysis on WHIP longitudinal data shows that the degree of Italian real wage procyclicality significantly increases after controlling for composition bias: this result is robust to several checks and it is consistent with Solon, Barsky and Parker’s 1994 seminal paper on the US. Finally, I discuss the effects of the 90’s labor market’s reforms on Italian real wage cyclical behavior.

1 INTRODUCTION

The goal of this work is to shed new light on real wage behavior over the business cycle in order to contribute to answer a very debated question in the economic literature: are real wages acyclical or procyclical over the business cycle?

The weak cyclical behavior showed by aggregate real wage statistics has been considered by the literature a salient feature of the business cycle. In the well-known work “Business Cycle Fluctuations in U.S. Macroeconomic Time Series” (1998), Stock and Watson found that real wages in the US. were weakly procyclical over the period 1947-1996 and weakly volatile, while employment - in terms of worked hours - was highly procyclical and had the same variability shown by the output.

Thus, real wages’ weak cyclical behavior and employment’s high procyclicality were considered stylized facts of labor market. In order to explain them, macroeconomists set up theoretical models in which employment fluctuations were the result of shifts of labor demand curve along a stable and highly elastic short-run labor supply curve, i.e. efficiency-wage models, insider-outsider models or implicit contracts models. In these models, only an *highly elastic short-run labor supply* curve could conjugate the weak procyclicality shown by real wages with employment’s high procyclicality. But an highly elastic labor supply is in contradiction with the majority of microeconomic evidence, pointing to a quite rigid labor supply, especially for the male component of the labor force.

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Beside those theories that predict acyclical real wages, there are theories that predict that real wages are actually procyclical over the business cycle: employment fluctuations are described as the result of shifts of labor demand curve along a positively sloped short-run labor supply curve. Labor demand shifts could be caused by productivity shocks (e.g. real business cycle models) or nominal disturbances (e.g. Keynesian models). The hypothesis of procyclical real wages has important implications on short-run labor supply's elasticity: if real wages are actually procyclical over the business cycle, there is no need for short-run labor supply curve to be highly elastic afterall. This would confirm the micro-level results.

An explanation can be found in Stockman (1983). He proved that aggregate real wage statistics - commonly used in the aggregate time series - show weak procyclicality over the business cycle because of a statistical bias. In particular, real wage aggregate statistics are constructed in a way that gives greater weight to low-skill workers during expansions than during recessions. So they exhibit a "*composition bias*" that would obscure - in a countercyclical direction - the true real wage procyclical behavior, as I discuss in Section 2.

Composition bias has been analyzed by many empirical studies, e.g. Bils (1985), Mitchell, Wallace and Warner (1985), Solon, Barsky and Parker (1994). The majority of these studies following the strategy proposed by Stockman found strong evidence of real wage procyclicality.

My aim is to analyze the behavior of Italian real wages over the business cycle by controlling for such a composition bias. Following the econometric strategy proposed by Solon, Barsky and Parker (1994), I estimate the cyclicity of Italian *real weekly wages* over the 1985-2003 period using *longitudinal data* from *WHIP*, the most suitable dataset for this kind of analysis (data issues are discussed in Section 4). I find that Italian real weekly wages are procyclical and their degree of procyclicality significantly increases once I control for the composition bias, suggesting that the aggregate real weekly wage statistics are contaminated by a downward bias. This finding - supported by several robustness checks - is consistent with Solon et al.(1994)'s results for the US. Despite the high procyclicality shown over the 1985-2003 period, Italian real wage response to business cycle fluctuations experiences a slight decline after 1997, due to the effects brought by the 90's labor market reforms.

In Section 2, I review the "*composition bias*" issue. Section 3 focuses on the econometric strategy I use to estimate Italian real wage cyclicity. Data, descriptive statistics and the sample selection procedure are presented in Section 4. Section 5 reports the estimation results - whose robustness is checked in Section 6 - and the comparison to the US. Section 7 focuses on the effects of the 90's labor market reforms on the cyclicity of Italian real wages. Finally, Section 8 discusses the main results and their implications and it concludes.

2 COMPOSITION BIAS

As pointed out by Stockman in 1983, the true procyclicality of real wages is obscured in aggregate time series because of a composition bias: to explain the point, I follow Solon et al.(1994). The aggregate wage statistics are constructed in a way that gives greater weight to low-skill workers during expansions than during recessions. In fact, the aggregate hourly real wage statistics are computed as the ratio of "*total wage bill* B_t " at time t , and "*total work hours* H_t ". Suppose that workers are "*divided into J groups, with $j = 1, 2, \dots, J$, with B_{jt} denoting the j -th group's wage bill and H_{jt} its total work hours*", so that $W_{jt} = \frac{B_{jt}}{H_{jt}}$ is the j -th group's average hourly earnings. Then, the overall aggregate wage statistic W_t can be written as:

$$W_t = \frac{B_t}{H_t} = \sum_{j=1}^J \frac{B_{jt}}{H_t} = \sum_{j=1}^J \frac{H_{jt}W_{jt}}{H_t} \quad \text{with } j = 1, \dots, J \quad (1)$$

Equation (1) shows that "*the aggregate wage statistic is a weighted average of the group-specific wage statistics with the groups weighted by their hour shares*".

If such aggregate average earnings are used to measure real wage cyclicity over the business cycle, one is implicitly assuming that the labor force composition stays constant over the business cycle. But this clearly does not happen. In fact, labor force composition - by age, sex and race - varies considerably with the business cycle. In particular, groups' hour shares vary with the business cycle. Several empirical studies¹ showed that work hours of low-wage groups (such as youth, black and less educated workers) are "*more cyclically variable than those of high-wage groups*". Thus, "*the aggregate wage statistics give greater weight to low-skill workers during expansions than during recessions*".

For instance², suppose that firms tend to lay off lower-skilled and/or less senior workers during the recession phases of the business cycle and to retain workers who are - on average - older/more skilled than the fired ones. The result is that the quality of the workers who are still employed will tend to increase. So, the average aggregate real wage will increase because of the exclusion of less-skilled groups from the labor force, even if no increase in the real wage per worker has occurred. Viceversa, during an expansion, the distribution of the labor force shifts towards unskilled/younger groups of workers, causing a decrease in the average real wage aggregate statistic, even if no worker had actually faced a cut in his wage. This induces a *countercyclical* composition bias. In conclusion, if this kind of statistics are used to estimate real wages cyclicity over the business cycle, the true procyclicality that the typical worker of each group actually faces will be underestimated.

Let's consider the cyclical variation in the aggregate wage statistic with respect to a cycle indicator Y_t , where Y_t is real GDP. Equation (1) can be rewritten in logarithms as:

$$\frac{d \ln W_t}{d \ln Y_t} = \sum_{j=1}^J \frac{H_{jt}}{H_t} \frac{W_{jt}}{W_t} \frac{d \ln W_{jt}}{d \ln Y_t} + \sum_{j=1}^J \frac{W_{jt}}{W_t} \frac{d (H_{jt}/H_t)}{d \ln Y_t} \quad (2)$$

¹Kosters and Welch (1972), Okun (1973), Clark and Summers (1981), Mitchell, Wallace and Warner (1985) and Kydland (1984).

²Mitchell, Wallace and Warner (1985).

The resulting measure of the aggregate real wage cyclicalities is a "*weighted average of the cyclical wage changes experienced by the J groups*" (the component that reflects the true wage cyclicalities) "*plus a second term reflecting the cyclical change in the skill composition of total work hours*". This last term represents the *composition bias*: "*if groups with low relative wages W_{jt}/W_t have procyclical hour shares, the second term contributes a countercyclical bias*".

How could the composition bias be corrected? Given that the measurement problem in aggregate wage data is caused by the cyclically shifting weights, a solution is to construct a wage statistic that gives fixed weights to the exact same workers over time. By following the same workers with fixed weights over time, labor force composition can be held constant over the business cycle. To do so, one needs to use *longitudinal micro-level data*. This is the strategy followed by Solon et al.(1994) to estimate US. real hourly wage cyclicalities over the 1967-1987 period. They find that after controlling for composition bias, US. real wages are actually *highly procyclical*, though aggregate real wage statistics have revealed - for the same period - weak procyclicalities.

Men and Women

Solon et al.(1994) estimate a different real wage cyclicalities for men and women, hence the aggregate real wage statistics are biased not only because of the composition bias described so far, but also because of the aggregation of women and men. Formally, let f and m denote women and men. Let β_{3m} and β_{3f} be "*the true wage cyclicalities for men and women*" and let δ be "*the proportional gap between wages paid for cyclically marginal hours of work and wages paid for nonmarginal hours*". "*If, within each gender, the cyclically marginal hours are less skilled than the nonmarginal hours, $\delta < 0$* ". Then, equation (2) can be rewritten³ as:

$$\begin{aligned} \frac{d \ln W_t}{d \ln Y_t} \cong & \frac{W_m}{W} \frac{H_m}{H} \beta_{3m} + \frac{W_f}{W} \frac{H_f}{H} \beta_{3f} + \left(\frac{W_f - W_m}{W} \right) \frac{H_m}{H} \frac{H_f}{H} \left(\frac{d \ln H_f}{d \ln Y} - \frac{d \ln H_m}{d \ln Y} \right) \\ & + \delta \left[\frac{W_m}{W} \frac{H_m}{H} \frac{d \ln H_m}{d \ln Y} + \frac{W_f}{W} \frac{H_f}{H} \frac{d \ln H_f}{d \ln Y} \right] \end{aligned} \quad (3)$$

Equation (3) shows that "*the cyclicalities of aggregate wage statistic is approximately a weighted average of the true cyclicalities of men's and women's wages plus two composition bias terms. The first of these reflects cyclical variation in the gender composition of total work hours. The second reflects cyclical variation in the skill composition of each gender's hours*".

If women are paid less than men and have less cyclically variable hours, the gender composition term by itself imposes a procyclical bias, while the skill composition bias from the last term is countercyclical (since $\delta < 0$). This is what Solon et al.(1994) found for US. data. In particular, they found that the procyclical gender composition bias is smaller than the countercyclical bias due to the last term - since the gender difference in hours cyclicalities is not very large. Therefore, the resulting aggregate wage statistics turn out to be downward biased. Moreover, Solon et al.(1994) found that real wages are less procyclical for women than for men, meaning that $|\beta_{3f}| < |\beta_{3m}|$, though the estimates for women are not statistically significantly different from zero.

³ All proofs in Solon et al.(1994).

3 ECONOMETRIC STRATEGY

To estimate Italian real wage cyclicity over the 1985-2003 period, I use the econometric strategy proposed by Solon et al.(1994). In particular, the following statistical model for characterizing the cyclicity in aggregate real wage data is used:

$$\ln W_t = \gamma_1 + \gamma_2 t + \gamma_3 t^2 + \gamma_4 (Y_t - \delta_1 - \delta_2 t - \delta_3 t^2) + \varepsilon_t \quad (4)$$

where W_t is the aggregate real wage statistic in year t , Y_t is real GDP - as business cycle indicator - and ε_t is a random error term. In particular, the wage measure I use in my analysis, W_t , is the real *weekly* wage, since long longitudinal data on hourly wages are not available in Italy. The implication of using weekly instead of hourly wage is discussed in the next section.

In order to analyze the cyclical components of wage and real GDP changes, a quadratic temporal trend is included in equation (4), and the business cycle indicator is expressed as a deviation from its own quadratic trend. Since ε_t is typically highly serially correlated and non stationary, by first-differentiating equation (4) one gets:

$$\Delta \ln W_t = \beta_1 + \beta_2 t + \beta_3 \Delta \ln Y_t + v_t \quad (5)$$

where:

$$\begin{aligned} v_t &= \Delta \varepsilon_t \\ \beta_1 &= \gamma_2 - \gamma_3 + \gamma_4 (\delta_3 - \delta_2) \\ \beta_2 &= 2 (\gamma_3 - \gamma_4 \delta_3) \\ \beta_3 &= \gamma_4 \end{aligned}$$

Equation (5) is estimated with ordinary least squares (OLS), so that $\hat{\beta}_3$ is the OLS estimate of the cyclicity of W_t with respect to the business cycle indicator: $\hat{\beta}_3 > 0$ if W_t is procyclical, $\hat{\beta}_3 = 0$ if W_t is acyclical and $\hat{\beta}_3 < 0$ if W_t is countercyclical. In the *first phase* of the analysis - the unbalanced sample - $\Delta \ln W_t$ is constructed by computing the log of the average real wage in year t , while in the *second phase* of the analysis, the dependent variable $\Delta \ln W_t$ is instead computed using the sample mean of the log real wage in year t among the workers of the balanced sample⁴. Results from OLS estimation of equation (5) are presented in Section 5, after data issues are discussed in the next section.

⁴As in Solon et al.(1994) for comparability.

4 DATA AND SAMPLE SELECTION

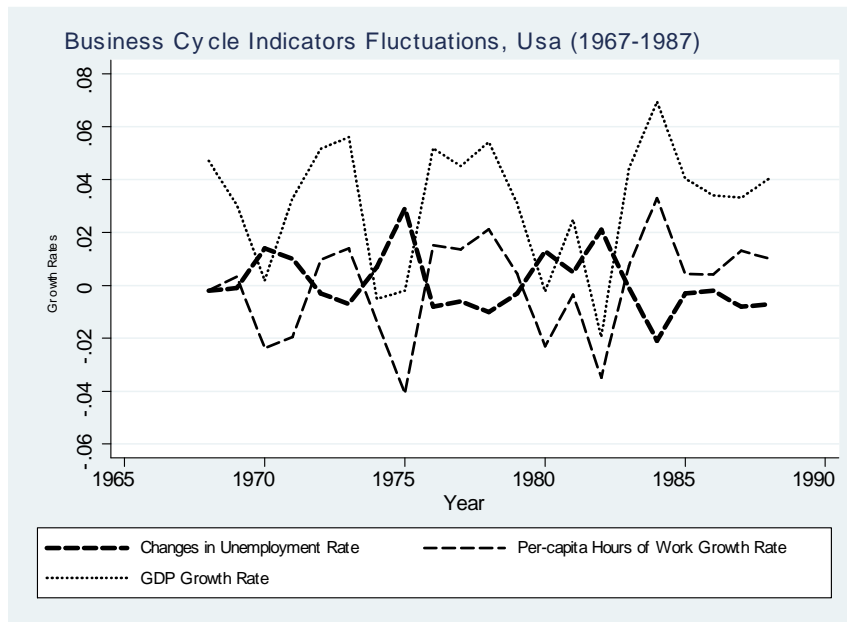
4.1 DATA AND DESCRIPTIVE STATISTICS

Business Cycle Indicator

The indicator of the stage of the business cycle I use is GDP at 2000 constant prices, as provided by Istat (Conti Nazionali 2008). Actually, Solon et al.(1994) use three different business cycle indicators to estimate real hourly wage cyclicality for the US.: the unemployment rate, real GNP (deflated by GNP implicit deflator) and per-capita hours of work (computed as the product of the employment/population ratio and the average annual work hours of the employed). This because the three indicators are highly correlated in the US. Figure 1 shows the three US. cycle indicators (in growth rates) from 1967 to 1987, i.e. the period analyzed by Solon et al.(1994). As one can notice from the graphs, the changes in the unemployment rate are highly negatively correlated with the GDP growth rate: the correlation coefficient between the two growth rates, ρ , is in fact equal to -0.86 . Per-capita hours of work shows a strong positive contemporaneous correlation with the output ($\rho = 0.89$). Now, let's look at Figure 2 - that depicts the three cycle indicators (in growth rates) over the 1985-2003 period in Italy. Differently from the US., Italian unemployment rate and per-capita hours of work⁵ are *weakly* correlated with the output, with $\rho = -0.13$ and $\rho = 0.35$, respectively. An explanation for this divergence can be found in institutional differences between the two countries, as several institutional constraints influence Italian change in unemployment and worked hours, such as employment protection laws and restrictions to overtime use. Thus, real GDP turns out to be a more proper indicator of the business cycle.

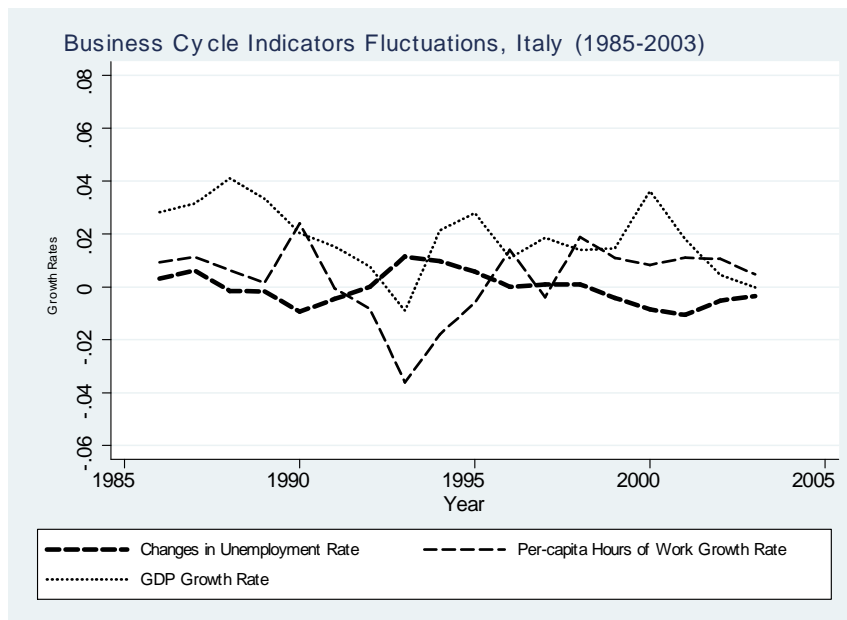
⁵Computed as in Solon et al.(1994). Source for both: Istat.

Figure 1. Business Cycle Indicators in the US., 1967-1987



Source: Economic Report of the President 2009, Table B-2, B-3, B-32, B-36, B-39,B-42.

Figure 2. Business Cycle Indicators in Italy, 1985-2003



Source: Conti Nazionali 2008 and Le ore lavorate per la produzione del Pil, Istat, 2008.

Wages from WHIP

The analysis of Italian real wage cyclicalities for the 1985-2003 period is based on panel data WHIP (Work Histories Italian Panel), an administrative dataset with information on the weekly wages of a large sample of Italian employees (about 100,000 yearly observations) in private firms. WHIP data include information not only about a certain number of characteristics of each worker (such as his/her annual labor income, work days and work weeks, gender, year of birth, geographical region where she/he works, work start/end dates, job qualification) but also about the firm where he/she works (such as industry and size). As the actual number of hours worked by an employee is not observable in WHIP, *hourly wages* cannot be computed. The number of "work days" and "work weeks" of each employee in each job spells and his/her total remuneration (before taxes) are however known, which allows us to compute the employee's daily (weekly) wage as the ratio of worker's annual remuneration and his/her worked days (weeks). Notice that this wage measure includes overtime work.

Alternative datasets that record Italian hourly wages are EU-SILC/ECHP and SHIW (Bank of Italy): both have shorter and more discontinuous longitudinal dimension and are not suitable for this kind of analysis. Moreover, WHIP data on wages are more reliable, as they are drawn from administrative archives. To the best of my knowledge, Italian real wage response to business cycle fluctuations is studied only by Peng and Siebert (2006), who use ECHP data to analyze Italian male real wage cyclicalities over the 1994-2001 period, distinguishing between job-stayers and movers and between Northern and Southern Italy.

The wage measure I use in my analysis is the real *weekly* wage, deflated by the consumer price index "Indice nazionale dei prezzi al consumo per l'intera collettività" (NIC senza tabacchi), a monthly price index that refers to the generality of consumption of domestic households, elaborated by Istat⁶. The main reason for not choosing real daily wage as a wage measure is the belief that the estimate of real wage cyclicalities could turn out to be biased. In fact, Contini, Filippi and Malpede (2000) observed that in the South of Italy, employers report a number of working days which is lower than the actual one - to the national social security archives (INPS). This Southern firms behavior leads to overestimate the resulting daily wage measure, with an higher overestimation during recessions than during booms.

The implication of using weekly instead of daily wages is that real weekly wages are expected to be more procyclical than daily wages: the reason can be found in the way they are computed. In fact, assuming that yearly earnings are held constant, it happens that work days vary more than work weeks over the business cycle. As real daily wages are constructed by dividing annual labor income by total worked days, it means that the denominator of daily wage statistic fluctuates more than the weekly wage's one - the numerator being constant. For the same reason real weekly wages are expected to be more procyclical than real hourly wages over the cycle. However, when I use real daily wage as alternative wage measure to check

⁶In Solon et al.(1994) real hourly wages are deflated by using the GNP implicit deflator. In my analysis, Italian real wages are deflated by using the CPI (NIC) instead of the GDP implicit deflator. By definition, the CPI reflects the prices of a representative basket of goods and services *purchased* by the consumers, whereas the GDP deflator reflects prices of all goods and services *produced* within the country. Since the ratio between the value of imports as a percentage of total domestic demand in Italy is double (about 11%) than in the US (about 6%, source: OECD I-O tables) and since the 70% of the NIC basket is represented by items associated to sectors with higher exposure to international trade, the CPI represents a more proper deflator for Italian prices than the GDP deflator.

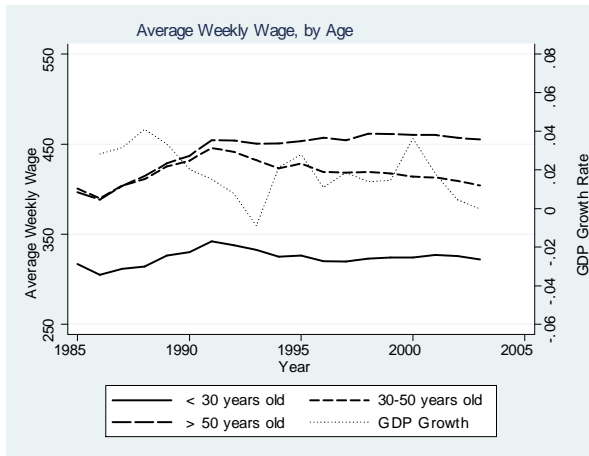
the robustness of the results⁷, the cyclicity shown by daily wages does not statistically significantly differ from that of weekly wages: because of the high correlation coefficient between work days and weeks, they show nearly coincident growth rates over the 1985-2003 period.

Figures 3 to 8 show WHIP weekly wage's sample averages by worker's age, gender, job qualification, geographical area, firm size and sector over the 1985-2003 period. Some regularities come out: in particular, real weekly wages are - on average - higher for men and for older workers; moreover, people who work in North-Western regions, in firms with more than 1,000 employees, and in service sectors earn more.

In each figure real GDP growth rate is reported, so that it is also possible to comment the cyclical behavior of real weekly wages over the period. After classifying workers by age, gender, job qualification, geographical area, firm's size and sector, no differences are visible in weekly wage cyclicity across groups. However, I will control for gender differences that might emerge, as discussed in Section 2, to let the data indicate whether they are relevant. In general, real weekly wages seem to be quite procyclical over the 1985-2003 period, though after 1993-94 the wage dynamics seem to be flatter than GDP's.

⁷In order to compare the two wage measures, the South of Italy is excluded from the sample to avoid potential measurement error, according to Contini, Filippi and Malpede (2000).

Fig.3. Average Weekly Wage by Age, 1985-03



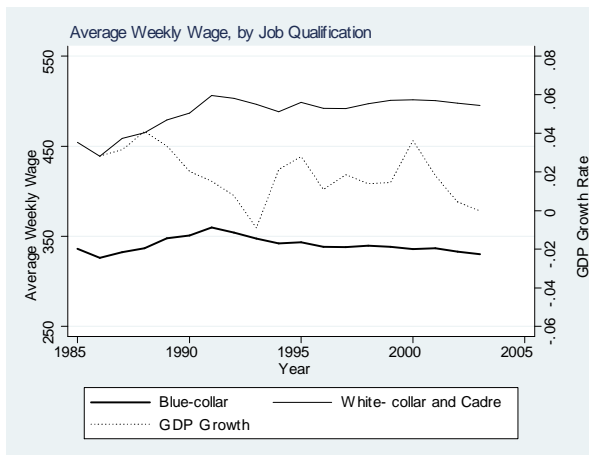
Source: WHIP

Fig.4. Average Weekly Wage by Gender, 1985-03



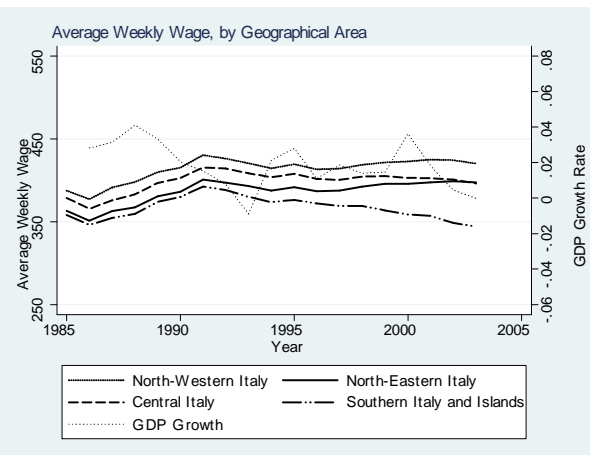
Source: WHIP

Fig.5. Average Weekly Wage by Job Title, 1985-03



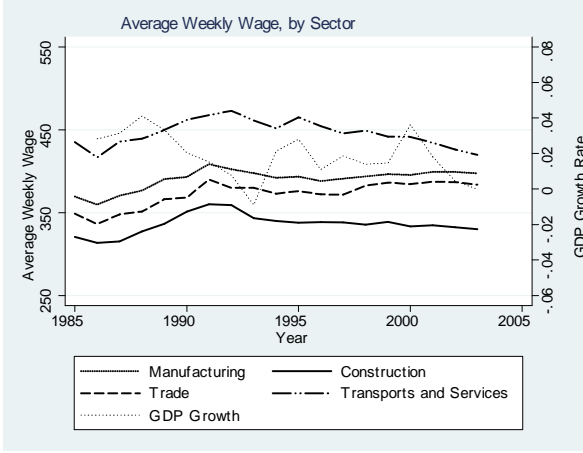
Source: WHIP

Fig.6. Average Weekly Wage by Area, 1985-03



Source: WHIP

Fig.7. Average Weekly Wage by Sector, 1985-03



Source: WHIP

Fig.8. Average Weekly Wage by Firm Size, 1985-03



Source: WHIP

4.2 WHIP SAMPLE SELECTION

Following Solon et al.(1994), a *two-step procedure* is followed to estimate the cyclicalities of real wages. As a first step, I estimate the aggregate real wage cyclicalities starting from an "unbalanced sample" that mimics aggregate national account statistics; then, in order to avoid composition bias, I construct a wage statistic based on a "balanced" sample. I expect real wages to be more procyclical using the balanced sample since the aggregate wage statistic is supposed to be affected by composition bias⁸.

Both the *unbalanced* and the *balanced* samples are constructed using WHIP longitudinal microdata. This approach differs from the one followed by Solon et al.(1994) since they compared the estimates of real hourly wage cyclicalities from BLS aggregate time series data to those from PSID longitudinal microdata. The fact that I use a *single* data source (instead of two) - and in particular WHIP longitudinal data - to construct both samples has a great advantage: measurement errors that potentially arise from the use of two different sources using different aggregation methods - are directly avoided.

Moreover, as discussed in Section 2, if women and men face different wage cyclicalities, the aggregate real wage statistic could also be affected by the bias caused by the aggregation of men and women. To verify it, I repeat the analysis for women, following Solon et al.(1994). Notice that we require women to respect the same selection criteria (discussed below) used to identify the male sample. This might imply that the female balanced sample is more selected than the male one, due to the presence of participation issues, but my work is not specifically devoted to explain women's behavior.

⁸To control for composition bias there are two alternative procedures: the first one is estimating equation (2) starting from a "balanced sample". The second possible procedure is estimating an equation in which one controls also for time-invariant individual characteristics of the workers (such as race, education). For example, Solon, Barsky and Parker estimated the following equation: $\Delta \ln W_{it} = \beta_1 + \beta_2 t + \beta_3 \Delta X_t + \beta_4 Y_{it} + v_{it}$, where Y_{it} is the worker's years of work experience in year t . Since we cannot observe such variables we will follow the first approach.

First Step: the Unbalanced Sample

The estimation of real wage cyclicality is based on repeated cross sections of white- and blue-collar workers⁹ employed in May (to avoid as much as possible patterns of seasonality in employment), without any further restriction. Outliers, identified as the 1st and the 99-th percentiles of the weekly wage's distribution, are removed from the sample. A sample of about 83,000 yearly observations is thus obtained. Tables 1 and 2 show some descriptive statistics for the WHIP 1998 sample: in particular, it can be noticed that in that year - on average - workers are 37 years old, they earn around 18,000 euros and they work 46 weeks a year. The 67% of the individuals from the sample are men, the 86% are prime-age workers and the 62% are blue-collarers. Moreover, the 60% of the sample workers work in manufacturing and more than the 60% in firms located in Northern regions of Italy. As expected, the majority of the workers is employed in small-size business firms (below 200 workers). With regard to the kind of worker's contracts, it must be said that the 93% of the cases are permanent worker and that only 10% are part-time workers.

Table 1. Descriptive statistics, 1998 unbalanced sample

Mean	Total	Men	Women
Age	37.09	37.86	35.53
Work days	270.91	281.76	249.23
Work weeks	46.56	48.40	42.88
Real annual income*	18,641	20,397	15,133
Real daily wage*	70.55	74.37	62.92
Real weekly wage*	393.83	416.00	349.53
Total	83,302	55,515	27,787

Source: WHIP; * Wages are deflated by NIC (2000 = 100)

⁹I drop "apprentices" and "managers" from the sample because of their peculiar wage behavior.

Table 2. Sample composition, 1998 unbalanced sample

Percentage (%)	Total	Men	Women		Total	Men	Women
<30 years old	26.74	24.03	32.17	<20	37.01	34.21	42.61
30-50 years old	59.27	60.18	57.45	20-199	31.01	31.58	29.90
>50 years old	13.99	15.79	10.38	200-999	14.03	14.64	12.81
				>=1000	17.94	19.58	14.68
Blue-collar	62.06	69.42	47.35				
White-collar and Cadre	37.94	30.58	52.65	North-Western	36.83	35.42	39.65
				North-Eastern	25.94	24.66	28.50
Manufacturing	51.99	53.90	48.18	Central Italy	19.23	18.92	19.86
Costruction	7.98	11.12	1.70	South e Islands	18.00	21.00	12.00
Trade	15.66	13.00	20.99				
Transports and services	24.37	21.98	29.13	Full-time job	90.53	97.31	77.01
				Part-time job	9.47	2.69	22.99
Cig*	5.25	6.39	2.97				
No cig	94.75	93.61	97.03	Permanent workers	93.50	93.77	92.96
				Temporary workers	6.50	6.23	7.04
Total	100	100	100	Total	100	100	100

Source: WHIP; * CIG ("Cassa Integrazione Guadagni") denotes whether the employee has received a wage supplement for temporary layoffs.

Second Step: the Balanced Sample

To avoid composition bias, I must have for each worker in the sample his/her wage observation in each year from 1985 to 2003. Thus, I restrict the sample to *prime-age men* - the group that is most likely to have positive work hours in every year of the sample period - who are born between 1944 and 1960. The birth year restriction assures that the sample members are between the ages of 25 and 59 throughout the sample period to minimize participation issues. Moreover, they are required to be white- and blue-collar workers, employed in May *every year* from 1985 to 2003 and to report at least 15 worked days in each year of the sample period. The resulting sample, made up of 8,936 men, is "balanced" in the sense that each year's wage information pertains to the exact same workers who meet all of the above criteria. The same procedure is followed to construct a female balanced sample, made up of 3,142 women. With respect to Solon et al.(1994), the size of these two samples is considerably greater, as in their work the male and the female balanced samples are made up of only 355 men and 146 women, respectively. Thus, in my analysis the accuracy of the estimates of the yearly means is increased. Moreover, let's notice that in Solon et al.(1994), no restriction on worker's job qualification, business sector, seasonality of employment has been made in constructing the "balanced sample": it is actually obtained by simply taking the same exact prime-age male (female) workers with positive labor income each year and at least 100 annual hours of work. Tables 3 and 4 report the descriptive statistics related to the male and the female balanced samples in 1998: after comparing them to Tables 1 and 2, one can notice that workers in the balanced samples - on average - earn and work more than those included in the unbalanced sample. In particular, the percentage of workers who are white-collars increases; they shift

from more mobile to less mobile sectors (from construction and commerce to manufacturing and services sectors) and from small to medium and large size firms. In terms of geographical area where they work, the percentage of workers in North-Western region increases, while that in the Mezzogiorno decreases. As expected, the percentage of workers who have a part-time contract and/or a temporary job sharply decreases.

Table 3. Descriptive statistics, 1998 balanced samples

Mean	Men	Women
Age	44.81	44.31
Work days	302.85	280.41
Work weeks	51.40	47.56
Real annual income*	36,185	27,143
Real daily wage*	95.19	79.28
Real weekly wage*	538.24	461.46
Total	8,936	3,142

Source: WHIP; *Wages are deflated by NIC (2000 = 100)

Table 4. Sample composition, 1998 balanced samples

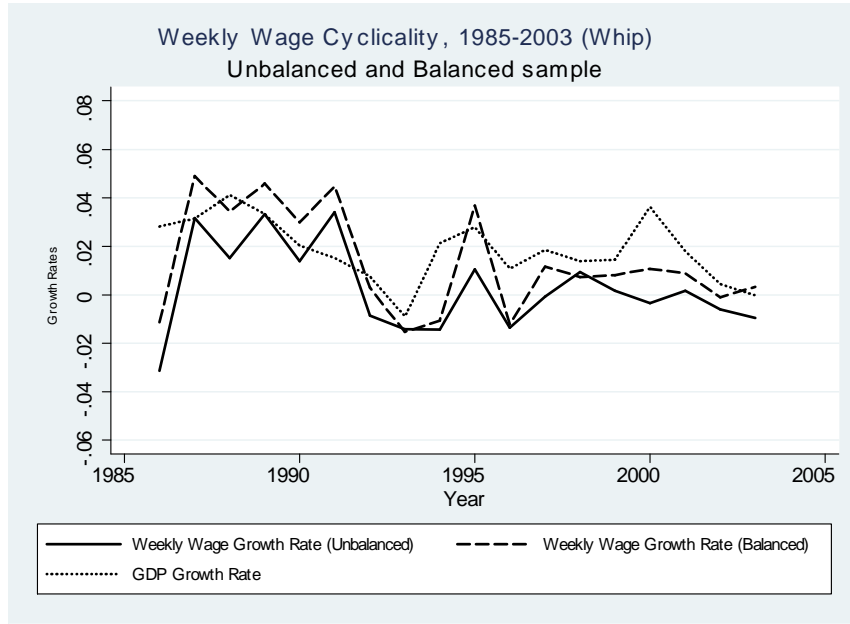
Percentage (%)	Men	Women		Men	Women
Men	100	-	<20	15.68	24.38
Women	-	100	20-199	27.19	28.90
<30 years old	-	-	200-999	19.87	17.38
30-50 years old	85.26	86.79	>=1000	37.26	29.34
>50 years old	14.74	13.21			
Blue-collar	55.00	33.42	North-Western	37.83	45.89
White-collar and Cadre	45.00	66.58	North-Eastern	22.96	25.30
			Central Italy	20.66	19.41
Manufacturing	57.63	53.91	South e Islands	18.55	9.39
Costruction	4.71	1.72			
Trade	9.33	17.09	Full-time job	99.43	83.45
Transports and services	28.33	27.28	Part-time job	0.57	16.55
Cig	5.12	4.33	Permanent workers	98.56	98.85
No cig	94.88	95.67	Temporary workers	1.44	1.15
Total	100	100	Total	100	100

Source: WHIP

5 ESTIMATION RESULTS

Figure 9 shows Italian weekly wages response to real GDP growth over the 1985-2003 period, where the wage statistic relates to the unbalanced and then to the male balanced sample: first of all, real weekly wages from WHIP seem to be procyclical over the period. Secondly, real weekly wages constructed on the balanced sample look like more procyclical than those constructed on the unbalanced sample, i.e. there is indication of an increase in real weekly wage procyclicality after one controls for composition bias.

Figure 9. Weekly Wage Cyclicity, Unbalanced and Male Balanced Sample, 1985-03



Source: WHIP

Table 5 shows the results from estimating equation (5), related to the two different phases of the analysis. The first column reports the estimated elasticity of real weekly wage with respect to GDP growth rate with the unbalanced sample: it refers to the sample that is supposed to be contaminated by composition and also by the bias that arises from aggregation of men and women. Regressing real weekly wage growth rate on real GDP growth rate, we have that the estimated coefficient of $\Delta \ln Y_t$ is equal to $\widehat{\beta}_3 = 0.474$, a statistically significant value implying that an increase in real GDP growth rate by one additional percentage point determines an increase in real weekly wage growth rate by about 0.5 percentage points. Therefore, even without controlling for composition bias, Italian aggregate real weekly wages are estimated to be *procyclical* over the 1985-2003 period. Since this estimate might be contaminated by the composition bias, let's control for it. The second column of Table 5 reports the results from estimating equation (5) on the male balanced sample. The estimated coefficient $\widehat{\beta}_{3m}$ is equal to 0.806, meaning that Italian real weekly wage are *highly* procyclical over the 1985-2003 period, confirming what Peng and Siebert (2006) found¹⁰. This value is nearly

¹⁰Though using different approach, data and methods, they found that after controlling for composition bias,

double compared to the estimates of $\widehat{\beta}_3 = 0.474$ for the unbalanced sample, representing the aggregate data subject to composition bias. To test if controlling for composition bias leads to a statistically significant increase in the estimated real weekly wage elasticity, a one-sided t-test is conducted in order to test the null hypothesis against the alternative $H_1 : \widehat{\beta}_{3m} > \widehat{\beta}_3$. With such test, the increase in weekly wage procyclicality is estimated to be significant at 0.07 significance level, suggesting that, over the 1985-2003 period, real weekly wages of Italian prime-age men who worked at least 15 days a year each year are considerably *more procyclical* than they appear in the aggregate wage statistics afflicted by composition bias.

Remark 1 *After controlling for composition bias, real weekly wages are significantly more procyclical than they are in the aggregate wage statistics.*

Now, let's look at the estimate for the female balanced sample, i.e. $\widehat{\beta}_{3f} = 0.507$; differently from Solon et al.(1994), I get that the estimate of real wage cyclicity for women is statistically significant. Thus, real weekly wages for women are procyclical, too: an increase of one additional percentage point in GDP growth rate leads to an increase of about 0.5 percentage points in their real weekly wage growth rates. After controlling for composition bias, the increase in real weekly wage procyclicality is not statistically significant: in fact, it is much smaller than that experienced by men. Therefore, if compared to men's, women's real weekly wages turn out to be less procyclical, confirming the results obtained by Solon et al.(1994), though here the gender difference in the estimated procyclicality is not statistically significant. This leads to think that if there is an aggregation bias in the aggregate weekly wage statistics, then it is actually very small. Despite what Solon et al.(1994) found for the US, I estimate that - over the 1985-2003 period - Italian women experience greater employment fluctuations: in particular, the cyclical variation in the total work weeks is estimated at 1.15 for women and only at 0.98 for men. The fact that over the cycle women experience greater employment fluctuations as well as lower wage fluctuations means that women's short-run labor supply is more elastic than men's. This might be an explanation for the gender difference in real wage cyclicity.

Table 5. Estimates of Real Weekly Wage Cyclicity in Italy, 1985-2003

	Unbalanced Sample	Male Balanced Sample	Female Balanced Sample
Cycle regressor: $\Delta \ln Y_t$			
	.474 ** (.217)	.806 *** (.251)	.507* (.262)
N=18			
R ²	0.18	0.36	0.18
Durbin-Watson	1.90	2.02	2.25

Sources: WHIP, Standard error estimates in brackets, *p<0.10,**p<0.05,***p<0.01

However, according to Solon et al.(1994), to obtain a bias-free measure of aggregate weekly wage cyclicity, I must estimate the weighted average of β_{3f} and β_{3m} - shown as the first two terms of equation (3). Based on the balanced samples, I found that the true wage cyclicity for women and men is $\widehat{\beta}_{3f} = 0.507$ and $\widehat{\beta}_{3m} = 0.806$, respectively. Moreover,

male "stayers in Northern Italy to have high cyclicity of real wages" over the 1994-2001 period, "higher in fact than the US and the UK".

from Table 1, I know that - in 1998 - $W_f/W = 0.89$ and $H_f/H = 0.31$, where H_f/H is the women's share of total work weeks, hence the women's share of the wage bill equals to $\frac{W_f}{W} \frac{H_f}{H} = 0.27$. According to equation (3), the true aggregate real wage cyclicalities is therefore estimated at $0.73 * (0.806) + 0.27 * (0.507) = 0.725$, though the unbalanced sample estimate displays a $d \ln W_t / dY_t$ of only 0.474. Once one controls for composition bias, the true real wage procyclicality is estimated to be about 0.7, meaning that an increase of one additional percentage point in real GDP growth rate leads to an increase of about 0.7 percentage points in real weekly wage elasticity. Therefore, Italian weekly wages seem to be highly procyclical¹¹.

The implied countercyclical composition bias in the aggregate wage statistic is therefore -0.251 , that can be decomposed into its two components: the cyclical change in the gender composition of total work weeks and the cyclical change in the skill composition of each gender's weeks. By substituting in $(W_f - W_m)/W = -0.17$, $\frac{d \ln H_f}{dY} = 1.15$ and $\frac{d \ln H_m}{dY} = 0.98$, the gender composition bias is estimated at -0.006 . As expected, since the gender difference in the estimated procyclicality is not statistically significant, the gender composition bias turns out to be very small. Let's notice that, despite what Solon et al.(1994) found, this term imposes a *countercyclical* bias as a consequence of the fact that, according to WHIP data, women are less paid than men and have more cyclically variable weeks over the 1985-2003 period. Calculated as a residual, the skill composition bias is estimated at -0.245 . The implied value of δ is -0.24 , denoting a 24 percent within-gender wage gap between cyclically marginally and nonmarginal weeks. In conclusion, over the 1985-2003 period, Italian real weekly wage statistics are affected by composition bias, which mainly arises from the cyclical change in the skill composition in each gender's work weeks.

Remark 2 *Over the 1985-2003 period, Italian weekly wages turn out to be affected by the bias arising from the changes in the skill composition of the labor force.*

5.1 COMPARISON TO US. ESTIMATES

To compare Italian real wage cyclicalities to the US.'s, I replicate Solon et al.(1994) work on US. aggregate data (BLS, Bureau of Labor Statistics) and then on longitudinal microdata (PSID, Panel Study of Income Dynamics), using both real *hourly* and *weekly* wages as wage measures over the 1976-1996 period¹². For comparison to my work, real GDP¹³ growth rate - instead of real GNP growth rate - is used as business cycle indicator.

Following Solon et al.(1994), the wage measure initially used is the average hourly (weekly) earnings of production and nonsupervisory workers on private nonfarm payroll industry¹⁴, generated by the BLS establishment survey. This wage measure includes work on overtime and second jobs and it is deflated by the implicit GDP deflator. Table 6 shows the estimated coefficients from OLS estimation of equation (5) of US. real hourly and weekly wage elasticity to real GDP growth rates over the 1976-96 period. The first two columns present results when

¹¹Using the unemployment rate and per-capita hours of work the main results are unchanged. Estimates available upon request.

¹²PSID data were collected annually through 1997, and biennially starting in 1999. In order to avoid manipulation of data for missing years (i.e. for 1998, 2000, 2002) I focus on the most recent and available years, that is from 1976 to 1996, that correspond to the 1977-1997 PSID interviews.

¹³Real GDP and implicit GDP deflator from *Economic Report of the President*, 2009, Table B-2 and B-3. GDP is at 2000 constant price.

¹⁴Average hours and earnings of production and nonsupervisory workers on private nonfarm payrolls by major industry sector, 1964 to date: *Employment and Earnings*, February 2009, Table B-2.

W_t is the BLS average hourly and weekly wage, respectively. According to BLS aggregate wage statistics, the hourly wage procyclicality is estimated at 0.155¹⁵, whereas the weekly wage procyclicality is estimated at 0.361, statistically significant values that denote a weak procyclical response of real wages to economic fluctuations.

Table 6. Estimates of Real Hourly and Weekly Wage Cyclicity in the US, 1976-96

	BLS-aggregated ¹		Male Balanced Sample ²		Female Balanced Sample ²	
Cycle regressor: $\Delta \ln Y_t$						
<i>Real Hourly Wage</i>	.155 ** (.068)		.568 ** (.238)		.320 (.267)	
<i>Real Weekly Wage</i>		.361 *** (.066)		.658 *** (.157)		.358 (.255)
N=20						
R ²	0.28	0.68	0.32	0.32	0.03	0.05
Durbin-Watson	0.87	0.92	2.55	2.77	3.25	2.99

Sources: ¹BLS and ²PSID. Male balanced sample: 255 men; female balanced sample: 156 women, *p<0.10, **p<0.05, ***p<0.01

Since these estimates are supposed to be contaminated by the composition bias, I proceed to estimate US. real wage elasticity to GDP growth starting from a male and a female balanced samples, drawn from the 1977-1997 cross-year family-individual level data files of the national longitudinal survey PSID. They collect yearly data on annual labor income, work hours and work weeks of each individual, which allow us to compute his hourly (weekly) wage as the ratio of his annual labor income to his annual work hours (weeks)¹⁶. Like the BLS average wage statistics, it includes work on overtime and second jobs and it is deflated by the implicit GDP deflator.

Like Solon et al.(1994), the male balanced sample is constructed by restricting the PSID data to men aged between 25 and 59 throughout the period, who were household heads every year of the sample period and who reported positive labor income and at least 100 hours of work each year: it contains 255 individuals, whereas the female balanced sample (made up of women who were household heads or spouses every year of the sample period and who are required to meet all the other above criteria) contains 156 individuals in each year of the 1976-96 period.

Based on the male balanced sample, I find that the real hourly and weekly wage elasticities to real GDP growth are respectively estimated at $\widehat{\beta}_{3m} = 0.568$ ¹⁷ and $\widehat{\beta}_{3m} = 0.658$, as shown by the third and the fourth columns of Table 6. If compared with the corresponding BLS estimates, the increase in real wage procyclicality is statistically significant with p-value < 0.01 after one controls for composition bias. Now, let's look at the last two columns of Table 6, related to the female balanced sample: for women, real hourly and weekly wage cyclicality are respectively estimated at $\widehat{\beta}_{3f} = 0.320$ and $\widehat{\beta}_{3f} = 0.358$. As the estimates for women are not statistically significant - confirming what Solon et al.(1994) found over the 1967-87 period, real wages are less procyclical for women than for men¹⁸. Notice that, as expected,

¹⁵Solon et al.(1994) estimated the hourly wage cyclicality at 0.293 (s.e. 0.077) for the 1967-87 period.

¹⁶Observations with "major assignments" imputed for labor income or work hours are excluded from the sample.

¹⁷Solon et al.(1994) estimated the hourly wage cyclicality at 0.617 (s.e. 0.165) for the 1967-87 period.

¹⁸For the 1967-87 period, I estimated real hourly wage elasticity to GDP growth to be: 0.323 (s.e.=0.120) from BLS and 0.629 (s.e.=0.158) from PSID male balanced sample. The estimated real weekly wage elasticity to

real weekly wage turns out to be more procyclical than hourly wage over the same business cycle: as already stated in Section 4, it might be due to the fact that over the cycle work hours are more cyclically variable than work weeks. Nevertheless, the greater procyclicality shown by real weekly wage is statistically significant (with p-value < 0.05) only for the BLS aggregate wage statistics, meaning that after controlling for composition bias the two wage measures give the same inferential results.

To get a free-bias measure of US. real wage elasticity to real GDP growth over the 1976-96 period, let's estimate the weighted average of β_{3f} and β_{3m} - shown as the first two terms of equation (3). From BLS, I know that the 1986 female wage-bill share is about $\frac{W_f}{W} * \frac{H_f}{H} = 0.34$ ¹⁹. Thus, the true real hourly wage cyclicity is estimated to be equal to 0.484, though the unbalanced sample estimate displays a $d \ln W_t / dY_t$ of only 0.155. Therefore, the resulting countercyclical composition bias in the aggregate wage statistic turns out to be estimated at -0.329 . With regard to the true real weekly wage procyclicality²⁰, it must be said that it is estimated around 0.560²¹, whereas the aggregate BLS wage statistics show that $d \ln W_t / dY_t$

GDP growth rate are: 0.539 (s.e.=0.122) from BLS and 0.697 (s.e.=0.165) from PSID male balanced sample. Hence, using real GDP instead of real GNP as business cycle indicator, the estimates are not significantly different from those obtained by Solon et al.(1994). Notice that the procyclicality shown by the aggregate real wages from BLS is significantly greater in 1967-87 than in 1976-96 period.

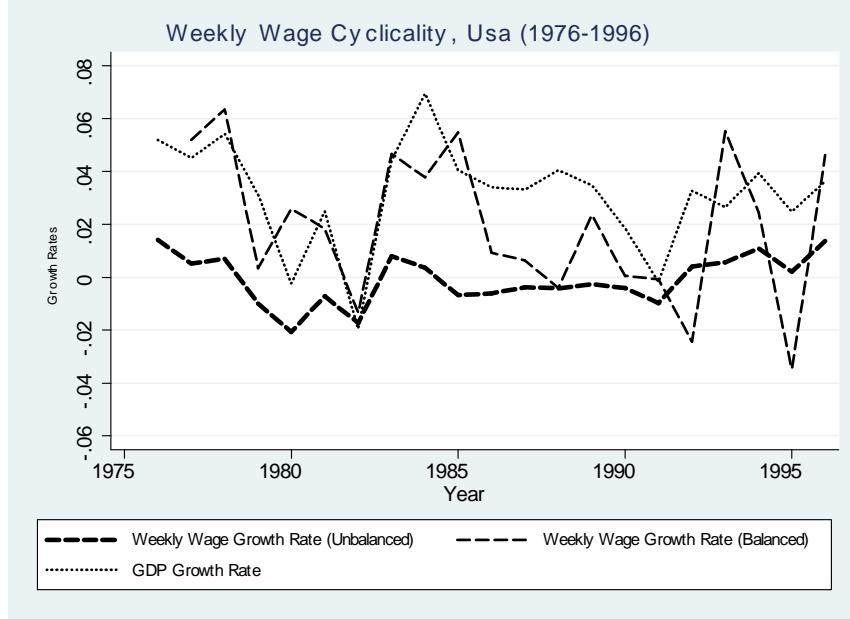
¹⁹Median hourly earnings of wage and salary workers paid hourly rates in constant (2000) dollars by sex and age, 1979-2000 annual averages: "Highlights of Women's Earnings in 2000", BLS, Report 952, Table 16. Employed persons by full- and part-time status and sex, 1970-2004 annual averages, Table 20: *Women in the Labor Force: a Databook*, BLS, May 2005, Report 985. Average weekly hours at work in all industries and in nonagricultural industries by sex, 1976-2004 annual averages, Table 21: *Women in the Labor Force: a Databook*, BLS, May 2005, Report 985.

²⁰Median usual weekly earnings of full-time wage and salary workers in constant (2000) dollars by sex and age, 1979-2000 annual averages, Table 13: "Highlights of Women's Earnings in 2000", BLS, Report 952.

²¹As data on number of worked weeks are not available from PSID, the true real weekly wage procyclicality is therefore computed by using the share of work hours as a proxy of the share of work weeks.

is equal to 0.361, implying a countercyclical composition bias of -0.199 .

Figure 10. Weekly Wage Cyclical, Unbalanced and Male Balanced Sample, 1976-96.



Source: Economic Report of the President 2009, Employment and Earnings, Feb. 2009, PSID.

Let's compare the estimates for Italian and US. real weekly wage response to economic fluctuations: if the second row of Table 6 is compared to Table 5, it turns out that, even if related to different business cycles, the cyclicalities of Italian and US. real weekly wage to economic fluctuations are of similar magnitude and not statistically different, though after controlling for composition bias, the increase in real weekly wage procyclicality experienced by men is greater in the US. Moreover, although real wages are less procyclical for women than for men in both countries, the gender difference in the real wage cyclical is statistically significant only with the US. estimates.

Remark 3 *Even if related to different business cycles, the estimates of Italian and US.'s real wage procyclicality are of similar magnitude and not statistically significantly different.*

Remark 4 *However, after controlling for composition bias, the US. estimates show a greater increase in real weekly wage procyclicality and a significant gender difference in real wage cyclical.*

6 ANALYSIS OF ROBUSTNESS

6.1 IMPOSING A RESTRICTION ON WORKED DAYS

In Solon et al.(1994)'s work, workers in the balanced sample were required to work at least 100 hours in each year of the 1967-1987 period. Here, as the actual number of hours worked by an employee is not observed, I impose a restriction on the number of actual worked days. So far, I have focused on workers who reported at least 15 work days.

What would happen to weekly wage cyclicalilty if the number of days worked by the individuals in the balanced sample was imposed to be greater?

Let's consider those who worked at least 100 days per year (the 75% of the female balanced sample and the 87% of the male balanced sample), then let's focus on those who worked at least 250 days per year (31% of female balanced sample and 46.5% of the male balanced sample) and finally those who worked at least 300 days per year (17% of the female balanced sample and 26% of the male balanced sample). Table 7 shows that real weekly wage procyclicality decreases when the number of required minimum work days increases. This result holds both for the male and the female balanced sample. Again, the greater procyclicality of men's real wage with respect to women's is confirmed to be not statistically significant.

Remark 5 *As the number of minimum required work days increases, real weekly wage procyclicality decreases.*

The reason might be the following: with such restrictions I am removing from the sample workers with higher level of job mobility, whose wage changes are larger than those within-job. One of the explanations for that is that employers use job-changes as a means of adjusting wages to the business cycle²². Therefore, what happens is that the estimated procyclicality of the remaining less mobile workers is therefore diminished.

Table 7. Estimates of Real Weekly Wage Cyclicalilty, Restriction on Worked Days, balanced samples,1985-2003

	Men Balanced Sample				Women Balanced Sample			
Cycle regressor: $\Delta \ln Y_t$								
<i>at least 15 days</i>	.806 *** (.251)				.507* (.262)			
<i>at least 100 days</i>		.799*** (.233)				.495 * (.261)		
<i>at least 250 days</i>			.773** (.236)				.468 ** (.215)	
<i>at least 300 days</i>				.691** (.271)				.345 (.243)
N=18								
R ²	0.36	0.35	0.34	0.30	0.18	0.19	0.19	0.13
Durbin-Watson	2.02	2.12	2.22	2.34	2.25	2.24	2.04	2.56

Sources: WHIP, Standard error estimates in brackets, *p<0.10,**p<0.05,***p<0.01

²²Devereux and Hart (2006).

7 STRUCTURAL BREAK

So far, I got that Italian real weekly wages have been highly procyclical over the 1985-2003 period. Nevertheless, the 1985-2003 period has been characterized by institutional reforms that might have affected Italian real wage response to economic fluctuations.

On the one side, some of these reforms were directly devoted to introduce greater flexibility to the Italian wage structure: the automatic mechanism (known as *scala mobile* system) through which Italian nominal wages were indexed to prices - was abolished in 1992. The purchasing power of wages is now set according to the Government's targeted rate of inflation, that, despite the aim of the reform, had the consequence of flattening out the wage dynamics²³.

In addition to that, a new bargaining system was introduced by the Income Policy Agreement (IPA) in July 1993. Beside the strongly centralized wage-setting bargaining arrangements - an additional bargaining level - regional or firm level - has been introduced; it is mainly devoted to the distribution of additional (top-up) components, i.e. wages over and above the industry-wide contractual wage, set according to firms' performance and local conditions. Despite the decentralized *top-up components* became more responsive to *local* unemployment after 1993²⁴, the power of central wage-setting in Italy has still remained high, the collective bargaining coverage being more than 80%.

On the other hand, later reforms were directly devoted to introduce greater flexibility to the "quantity" side of the labor market: new non-standard contracts (free lance/quasi subordinate & temporary agency workers jobs) were introduced in 1997 with the so-called *Pacchetto Treu* and in 2001 there was the liberalization of fixed term contracts. Leombruni (2008) found that after 1997 there has been a sharp increase in the Italian *gwt* (gross worker turnover) after the introduction of the new contracts brought by the *Pacchetto Treu*-reform.

Here, my aim is to verify if Italian real wage cyclicity has been affected by these institutional changes. A test has been conducted in order to determine if there was a structural break in Italian real weekly wage cyclicity over the 1985-2003 period.

In particular, I run the following specification:

$$\Delta \ln W_t = \alpha_3 + \beta_3 t + \gamma_3 \Delta \ln Y_t + \alpha_3' d + \beta_3' dt + \gamma_3' d \Delta \ln Y_t + v_t \quad (6)$$

where d is a dummy variable equal to 0 for period 1 and equal to 1 for period 2. The model is equivalent to estimating two separate models for the two periods.

Then, the coefficients α_3' , β_3' and γ_3' are tested to be jointly equal to zero: the resulting test statistic is distributed as a $F(k, N_1 + N_2 - 2k)$, where k is the number of the estimated parameters ($k = 3$ in our case), N_1 and N_2 are the number of observations in the two periods. Rejection of the null hypothesis means that the two periods do not share the same intercept, slope and trend, so there is a structural break in real weekly wage cyclicity.

Different alternatives have been tested on the unbalanced sample and the null hypothesis of non existence of a structural break in the time series is rejected when I check for the period before and after the year 1997. However, when the same test is run on the male balanced sample, I find that there is *weak* evidence of a structural break in 1997.

Table 8 presents the results from estimating equation (6) on the unbalanced and the male balanced samples when $d = 0$ for the period before 1997 (and $d = 1$ for the period after 1997).

²³Devicienti, Maida and Pacelli (2008)

²⁴Devicienti, Maida and Pacelli (2008).

Table 8. Estimates of Structural Break in Weekly Wage Cyclicity, Unbalanced and Male Balanced sample

	Unbalanced Sample	Male Balanced Sample
	1997	1997
<i>Regressors:</i>		
t	-.000 (.002)	-.001 (.002)
$\Delta \ln Y_t$.645 ** (.266)	1.061 *** (.311)
d	.065 ** (.029)	.013 (.027)
dt	-.003 (.002)	-.000 (.003)
$d\Delta \ln Y_t$	-.750 ** (.284)	-.831 ** (.317)
<i>constant</i>	-.005 (.026)	-.000 (.026)
Test $H_0:d,dt,d\Delta \ln Y_t=0$		
F(3,12)	5.15	3.18
Prob>F	0.016	0.063

Sources: WHIP, Standard error estimates in brackets, *p<0.10, **p<0.05, ***p<0.01

Let's look at the first column of Table 8, that relates to the unbalanced sample: the result of the test coefficients against 0 in the pooled specification is an F-test statistic equal to 5.15, distributed as $F(3, 12)$, with critical value 3.49 at 5% significance level. There is a statistically significant reduction in Italian real wage cyclicity after flexibility on the "quantity" side of the labor market has been increased by the Treu's reform: before 1997, the estimated real weekly wage procyclicality is 0.645, whereas, after that year, it turns out to be not significantly different from zero.

Why do I find a "reduction" in real weekly wage cyclicity after 1997? As already stated, it has been empirically found that the introduction of new-flexible contracts in 1997 led to an increase in Italian *gwt*. Thus, it might be the case that - as a consequence of the increase in the flexibility on the "quantity" side of the labor market - the flexibility on the "price" side has diminished. The 1997 reform might have given a new opportunity to firms to recur to these new contracts to adjust their labor demand needs. If this is true, they could have affected the way in which Italian labor market adjusts to economic fluctuations: it is reasonable to think that now, labor market adjustments could more often/easily occur through "quantity" rather than through "price" (i.e. wage) adjustments. Hence, the reduction in real wage elasticity could be due to the fact that after 1997 business cycle fluctuations affect - first of all - the quantity side. Moreover, we know that after 1993, the 80% of labor contracts are set according to the Government's targeted rate of inflation, thus inducing a general real wage flattening (as also documented in Figure 9). Though a structural break in 1993 has not been detected (suggesting that in my analysis the 1993 reforms do not have, *per se*, a measurable effect on real wage cyclical behavior), the sum of the effects brought by the 1993 and by the Treu's reforms might be responsible for the reduction in Italian real wage dynamics after 1997.

However, after controlling for composition bias the existence of a significant structural break in 1997 is found to be *weaker*. As the second column of Table 8 shows, the estimated real weekly wage procyclicality decreases from 1.061 to 0.230, a statistically significant value

that denotes mild procyclicality. The result of the test is equal to $F(3, 12) = 3.18$, with critical value 2.60 at 10% significance level. Hence, balanced sample workers experienced a significant but weaker structural break in their real weekly wage cyclicity. The reason might be the following: differently from those in the unbalanced sample, balanced sample workers turn out not to be affected by the 1997 introduction of new non-standard contracts as they are long-tenured workers, by construction. In fact, after the Treu's reform, the presence of *temporary workers* in the male balanced sample is lower (1.4% in 1998) than in the unbalanced sample (6.50% in 1998), where it starts increasing over time. Therefore, the slight decline in the wage cyclicity of balanced sample workers might be simply due to the effects of the 1993 reforms.

8 CONCLUSIONS

In this work, I estimate Italian real *weekly wage* cyclicalities over the 1985-2003 period, by using longitudinal data from WHIP. To do so, as discussed in Section 2, I followed the strategy proposed by Stockman (1983), who proved that the aggregate real wage statistics - commonly used to measure real wage cyclicalities - are affected by a "composition bias", i.e. a bias that might arise from the aggregation of low and high-wage groups of workers, experiencing different hours share's cyclicalities over the business cycle. In fact, they give more weight to low-skilled workers during expansion than during recession, thus inducing a countercyclical bias in the estimate of real wage's cyclicalities. After controlling for composition bias, real wages turn out to be considerably more procyclical than indicated by aggregate time series, that show near-noncyclicalities for the same period. Moreover, the true procyclicalities of real wages might be underestimated by the aggregation of women and men, who experienced different wage cyclicalities, as pointed out by Solon, Barsky and Parker in 1994. The findings of real wage procyclicalities has very important implications on the extent of short-run labor-supply curve's elasticity: if real wages are actually procyclical over the business cycle, the resulting estimated elasticity of the labor-supply curve is therefore diminished, confirming micro-level results.

As shown by Table 5, I find that in Italy, real weekly wages are procyclical over the 1985-2003 period: an increase of one additional percentage point in real GDP growth rate leads to an increase of about 0.47 percentage points in real weekly wage elasticity. However, after controlling for composition bias, real weekly wages are considerably *more procyclical* than they appear in the aggregate wage statistics afflicted by composition bias. In fact, the estimated procyclicalities becomes equal to 0.80: the increase in real wage procyclicalities is estimated to be significant at 0.07 significance level.

These results are consistent with Solon et al.(1994)'s findings for the US: they estimate that, after controlling for composition bias, real *hourly wage* cyclicalities has increased from 0.29 up to 0.62 over the 1967-87 period. In order to compare US. to Italian estimates, I replicate their work by using real weekly wage as a wage measure. As Table 6 shows, after controlling for composition bias, US. weekly wage procyclicalities is estimated to be about 0.66 for 1976-96 period, whereas the aggregate wage statistic displays a real wage elasticity of only 0.36. Though after controlling for composition bias the increase in real weekly wage cyclicalities is higher in the US. than in Italy, Italian and US. real weekly wage responses to economic fluctuations turn out to be of similar magnitude and not statistically significantly different. Secondly, I find real weekly wages to be significantly more procyclical for men than for women only in the US.

Moreover, as discussed in Section 6, the result of high procyclicalities shown by Italian real weekly wages is robust to differences in the number of minimum days worked by the individuals. I also find that real weekly wage procyclicalities experienced by men and women decreases when I focus only on those who work "hard" (i.e. almost 300 days a year): they are presumably workers with a low level of job mobility, whose wages therefore fluctuate less than those of job-movers.

Even if real weekly wages turn out to be strongly procyclical over the 1985-2003 period, a statistically significant reduction in Italian real weekly wage procyclicalities is detected in 1997, when testing for structural breaks in the time series. It might be interpreted as the consequence of two overlapping effects: on the one hand, after 1993, wages at the national

bargaining level are set according to the Government's targeted rate of inflation, with the consequence of flattening out the wage dynamics. On the other hand, the 1997 introduction of new non-standard labor contracts (free lance/quasi subordinate & temporary agency workers jobs) brought by the *Pacchetto Treu* led to an increase in the flexibility of the "quantity" side of the labor market, thus presumably reducing the flexibility on the "price" side. If conditioned to balanced sample workers, the 1997 structural break in real weekly wage procyclicality is significant but weaker: focusing on long-tenured workers, the effect of the Treu's reform disappears as those with new non-standard contracts are not included in this sample, by construction.

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